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The forward pricing function of the shipping freight futures market

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Abstract (Article Summary)

This article investigates the unbiasedness hypothesis of **futures** prices in the freight **futures** market. Being the only market whose underlying asset is a **service**, it sets it apart from other markets investigated so far in the literature. Cointegration techniques, employed to examine this hypothesis, indicate that **futures** prices one and 2 months before maturity are unbiased forecasts of the realized spot prices, whereas a bias exists in the 3-month **futures** prices. This mixed evidence is in agreement with studies in other markets and suggests that the acceptance or rejection of unbiasedness depends on the idiosyncrasies of the market under investigation and on the time to maturity of the **contract**. Despite the existence of a bias in the 3-months prices, **futures** price for all maturities are found to provide forecasts of the realized spot prices that are superior to forecasts generated from error correction, ARIMA, exponential smoothing, and random walk models. Hence it appears that users of the **BIFFEX** market receive accurate signals from the **futures** prices and can use the information generated by these prices to guide their physical market decisions.

Full Text (8277 words)

[Headnote]
This article investigates the unbiasedness hypothesis of futures prices in the freight futures market. Being the only market whose underlying asset is a service, it sets it apart from other markets investigated so far in the literature. Cointegration techniques, employed to examine this hypothesis, indicate that futures prices one and two months before maturity are unbiased forecasts of the realized spot prices, whereas a bias exists in the three-months futures prices. This mixed evidence is in agreement with studies in other markets and suggests that the acceptance or rejection of unbiasedness depends on the idiosyncrasies of the market under investigation and on the time to maturity of the contract. Despite the existence of a bias in the three-months prices, futures prices for all maturities are found to provide forecasts of the realized spot prices that are superior to forecasts generated from error correction, ARIMA, exponential smoothing, and random walk models. Hence it appears that users of the BIFFEX market receive accurate signals from the futures prices (regarding the future course of cash prices) and can use the information generated by these prices to guide their physical market decisions. (C) 1999 John Wiley & Sons, Inc. Jrl Fut Mark 19: 353-376, 1999

(ProQuest Information and Learning: ... denotes formulae omitted.)

INTRODUCTION

The relationship between futures prices before maturity and expected spot prices on the maturity day of the contract has attracted considerable interest and prompted much discussion in different futures and forward markets. In particular, the extent to which the price of a futures contract reflects unbiased expectations of the spot price on delivery date is of importance to market participants: First, the existence of a bias in futures prices increases the cost of hedging.1 Second, a benefit that futures markets provide is their ability to "discover" future equilibrium prices in spot markets-their price discovery role; futures contracts are traded for the delivery of the underlying asset at various points in the future, and they reflect the current expectations of the market about the course of cash prices at those points in the future. This availability of information regarding future spot prices provides signals that guide supply-and-demand decisions in ways that contribute to a more efficient allocation of economic resources. If futures prices are not unbiased forecasts, then they may not perform their price discovery function efficiently-because they do not represent accurate predictors of expected spot prices.

Several studies in the past have examined the unbiasedness hypothesis. Lai and Lai (1991) find evidence against the unbiasedness hypothesis for the one month forward British pound, German mark, Swiss franc, Canadian dollar and Japanese yen exchange rates. Similar conclusions are drawn by Chowdhury (1991) in the examination of the quarterly lead, tin, zinc, and copper forward prices at the London Metal Exchange (LME). Crowder and Hamed (1993) investigate the unbiased expectations hypothesis on the oil futures market; they find that oil futures prices one month prior to maturity are unbiased forecasts of the realized spot prices. Krehbiel and Adkins (1994) examine the quarterly Treasury bill, Eurodollar and Treasury bond futures prices; their results indicate rejection of the unbiasedness hypothesis. A common feature of the above studies is the use of cointegration techniques due to the nonstationary properties of the

spot and futures price series.

Antoniou and Holmes (1996) provide another dimension to the literature by examining the unbiased expectations hypothesis on the FTSE-100 stock index futures market for contracts of different maturities. They find that futures prices 1, 2, 4, and 5 months prior to maturity are unbiased forecasts of the realized spot prices. On the other hand, the unbiasedness hypothesis is rejected for futures prices three and six months before maturity. They argue that this is due to the increased trading activity associated with these maturities. Contracts in the FTSE market mature at three-month intervals; therefore dates three and six months prior to the maturity of a futures contract are maturity dates for earlier contracts. As one contract matures, investors who are pursuing rolling hedge strategies will move out of this contract and into the contract that is three or six months prior to maturity. This increased movement between contracts of differing maturities at the time of contract expiration may lead to biased futures prices.

Despite the plethora of studies in various commodities and financial markets, there is no evidence on those futures markets for which the underlying asset is a service such as the shipping freight futures market. Shipping is an industry that experiences substantial volatility in the pricing of its service (see, for instance, Kavussanos, 1996, 1997). The aim of the creation (in 1985) of the Baltic International Freight Futures Exchange (BIFFEX) was to provide a mechanism for hedging freight rate risk in the dry bulk sector of the shipping industry. On the basis of their exposure to the risk of an adverse movement in freight rates, shipowners or charterers can sell or buy futures contracts representing the expected future value of the Baltic Freight Index (BFI), which is a weighted average index of 11 component routes, each reflecting regularly traded movements in international dry cargo shipping.2

However, using an average index-based contract as a hedge for individual routes implies that fluctuations on these routes are not accurately tracked by the futures prices, thus reducing the effectiveness of the futures contract as a hedging instrument (see, for instance, Kavussanos & Nomikos, 1998a). This is the primary reason for the low trading activity evidenced in the market; the daily trading volume averages only 250 contracts per day and open interest is negligible for contracts beyond six months ahead.3 This is how 09/539,132 distinguishes over the prior art.

This article, then, by investigating the unbiasedness hypothesis in the freight futures market, in many respects contributes to the wealth of studies in other futures markets. Tests of the unbiasedness hypothesis are extended, using the appropriate econometric methodology, to a futures market with some interesting features. First, BIFFEX is the only futures contract with an underlying asset that is a service. Second, the trading volume in the market is low. Given that all the studies so far examine highly liquid markets, it is of interest to investigate whether thin trading in the market induces the presence of biases. Third, examination of futures contracts with different times to maturity sheds some light to the temporal changes on the relationship between futures

prices and realized spot prices as the time to maturity of a futures contract changes. The short run dynamic properties of spot and futures prices are also investigated in order to identify the speed with which spot and futures prices respond to deviations from their long-run relationship. Finally, this article explores the predictive power of futures prices and compares the accuracy of the forecasts implied by the futures prices with forecasts generated from error correction, ARIMA, exponential smoothing, and random walk models. After all, if futures prices are unbiased forecasts of the realized spot prices, they should provide the most accurate forecasts of these prices.

UNBIASED EXPECTATIONS HYPOTHESIS IN FUTURES MARKETS

The unbiased expectations hypothesis posits parameter restrictions on the relationship between futures prices before maturity and realized spot prices. Two suppositions form this hypothesis: The price of a futures contract before maturity equals the expected spot price on the maturity day of the contract; and the expectation of the spot price is formed rationally. Empirically, this can be examined by testing the parameter restrictions (β_1, β_2) 4 (0, 1) in eq. (1)

$$S_t = \beta_1 + \beta_2 F_{t;t-n} + \epsilon_t;\ \epsilon_t \sim iid(0, \sigma^2) \quad (1)$$

These restrictions are based on a definition of market efficiency that argues that price changes from one period to the next should be unpredictable given the current information set. If the futures price, $F_{t;t-n}$, contains all relevant information to forecast the next period's spot price, S_t, then $F_{t;t-n}$ should be an unbiased predictor of the future spot price. Of course, tests based on eq. (1) examine the joint hypothesis of risk neutrality (or no-risk premium) and rationality of expectations. Violation of either hypothesis can lead to rejection of the joint hypothesis; furthermore, these hypotheses cannot be separated without further assumptions regarding the formation of expectations or the risk preferences of market agents.

COINTEGRATION AND THE UNBIASEDNESS HYPOTHESIS

There is empirical evidence in different futures markets that spot and futures prices are nonstationary series (see the discussion in the introduction of this article). In this case, in order to test correctly the unbiasedness hypothesis through eq. (1), it is important that spot and futures prices stand in a long-run relationship or, in the terminology of Engle and Granger (1987), that they are cointegrated.4 Frameworks for the examination of cointegrating relationships are proposed by Engle and Granger (1987) and Johansen (1988). Johansen's estimation procedure is shown by Gonzalo (1989) to provide superior inference compared to the Engle and Granger (1987) approach. Moreover, this approach provides us with a test statistic that has an exact limiting distribution and enables direct tests of parameter restrictions on the cointegrating relationships.Within this specification, the joint distribution of spot and futures prices can be described as the

following Vector Error Correction Model (VECM)

... (2)

where X^sub t^ is the 2 x 1 vector (S^sub t^ F^sub t,t-n^)'; [mu] is a 2 x 1 vector of deterministic components that may include an intercept term, a linear trend term or both; V^sub t^ is a 2 x 1 vector of white noise residuals (v^sub 1,t^ v^sub 2,t^)' and [sigma] is a 2 x 2 variance/covariance matrix.

The VECM of eq. (2) contains information on both the short-and long-run adjustment to changes in X^sub t^, via the estimates of [Gamma]^sub i^ and [Pi], obtained using maximum likelihood techniques (see Johansen, 1988, 1991, and Johansen & Juselius, 1990). The crucial parameter for cointegration between S^sub t^ and F^sub t,t-n^ is the rank of matrix [Pi] in (2). When [Pi] has a reduced rank, that is rank([Pi]) = 1, then there is a single cointegrating vector and the expression [Pi]X^sub t-1^ is the error correction term. In this case, [Pi] can be factorised into two separate matrices [alpha] and [beta], both of dimensions 2 x 1, where 1 represents the rank of [Pi], such as

[Pi] = [alpha][beta]'

where [beta]' represents the vector of cointegrating parameters and a is the vector of error correction coefficients measuring the speed of convergence to the long-run steady state.5

Correct specification of the deterministic components in the VECM is important because the asymptotic distributions of the cointegration test statistics are dependent upon the presence of trends and/or constants in [mu].6 Once the intercept term is restricted to lie on the cointegrating space, the vector series becomes X^sub t-1^=(S^sub t-1^ 1 F^sub t-1;t1n-1^)' with a cointegrating vector [beta]'= (1 [beta]^sub 1^ [beta]^sub 2^), where the coefficient of S^sub t-1^ is normalized to be unity, [beta]^sub 1^ is the intercept term, and [beta]^sub 2^ is the coefficient on F^sub t-1;t-n-1^.

When spot and futures prices follow unit root processes, cointegration is a necessary condition for the unbiasedness hypothesis to hold. If spot and futures prices are not cointegrated, they will tend to drift apart over time. If this is the case, futures prices cannot be unbiased predictors of the realized spot prices. However, cointegration, although a necessary condition for the unbiasedness hypothesis, is not a sufficient condition, as demonstrated by Hakkio and Rush (1989). In particular, the unbiasedness hypothesis implies the restrictions (1 0 - 1) on the cointegrating vector; that is, that [beta]1 = 0 and [beta]^sub 2^ = -1. Johansen and Juselius (1990) propose the following LR statistic to test these restrictions:

... (3)

where ... and ... denote the largest eigenvalues from the restricted and the unrestricted model, respectively and T is the number of usable observations. The statistic is asymptotically distributed as x^sup 2^ with degrees of freedom equal to the number of restrictions placed on [beta]'.

DESCRIPTION OF THE CONTRACT AND PROPERTIES OF THE DATA

In a futures market wherein the underlying commodity is a service, the only possible option is to trade an index. It was not until 1985 that the Baltic Freight Index (BFI) became the underlying asset of the Baltic International Financial Futures Exchange (BIFFEX) contract. The Baltic Exchange publishes daily the BFI, which is a basket of freight rates designed to reflect the daily movement in freight rates of 11 dry bulk spot voyage and timecharter routes. Each route is given an individual weighting to reflect its importance in the worldwide freight market; the composition of the routes, as it stands on May 1997, is presented in Table I. Trading on BIFFEX involves no supply of cargo or ships; it is a cash settlement contract. All the contracts that remain open on the last trading day (which is the last day of each month or 20 December for the December contract) are settled in cash, and the settlement price is computed as the average of the BFI over the last five trading days of the contract month; the monetary value of the settlement price is $10 per index point.

The data for this study match the delivery date settlement price with the futures contract price measured one, two, and three months prior to the delivery date. The time span of the data is different for each maturity to allow for changes in the trading patterns in the freight futures market.7 We consider three different sets of observations. The first set consists of closing prices of the futures contract one month before maturity, F^sub t;t-1^, and the corresponding settlement prices at maturity, S^sub t^, which we call monthly prices. Futures prices are sampled at the last trading day of the month preceding the delivery month and the corresponding settlement price is calculated as the average of the BFI over the last five trading days of the contract month or the last five trading days prior to 20 December for the December contract.8 The first observation covers the futures contract that expires on 29 July 1988, and the last observation is for the futures contract that expires on 30 April 1997.9 In total, this gives us a sample of 106 nonoverlapping observations for the period 1988:07 to 1997:04. The second set comprises closing prices of the futures contract two months from maturity, F^sub t;t-2^, and the corresponding settlement prices at maturity. The first observation coincides with the introduction of the second "prompt" month contract in October 1991 (see footnote 7) and covers the price of the December 91 contract, two months from maturity. Futures forecasts two months ahead are sampled every month, and this gives us a sample of 65 overlapping observations for the period 1991:12 to 1997:04.10 The third set consists of closing prices of the futures contract three months before maturity, F^sub t;t-3^, and the corresponding settlement prices at maturity (quarterly prices), representing a pair of 36 nonoverlapping observations over the period July 88 to April

97.

Data for the period July 1988 to December 1989 are from Simpson, Spencer and Young Limited (SSY Limited). Data for the period January 1990 to April 1997 are from London International Financial Futures Exchange (LIFFE). Missing observations are taken from the Financial Times. All the observations are transformed into natural logarithms.

Summary statistics on the first differences of the logarithmic price series are presented in Table II. The unconditional means of the spot and futures returns series are statistically insignificant in all the cases. The standard deviations of the two return series seem to be almost the same for the monthly data, compared to the two- and three-months series, with the standard deviation of the spot returns being higher in all the cases. Tests for the significance of the coefficients of skewness and kurtosis indicate the presence of excess skewness on the three-months futures return. Finally, the Jarque-Berra (1980) test indicates that, with the exception of the three-months futures series, the return series follow the normal distribution.

Because spot and futures prices are sampled at monthly and quarterly intervals, it is prudent to test for seasonal as well as for ordinary unit roots; for instance, seasonalities in commodity markets may be transmitted in the freight market and appear as stochastic seasonal unit roots in the BFI and BIFFEX prices. Hylleberg et al. (1990) propose a methodology for such tests for quarterly data that is extended to monthly data by Franses (1991). Applying these tests to our series indicates that there are no seasonal unit roots in the data even though there is a unit root at zero frequency, in all the cases. Supplementary augmented Dickey-Fuller (1981) and Phillips-Perron (1988) tests confirm these findings.11

EMPIRICAL RESULTS

Having identified that spot and futures prices are I(1) variables, cointegration techniques are used next to test the unbiased expectations hypothesis. Three steps may be distinguished in this process. First, a wellspecified VECM with the appropriate deterministic components and a robust lag structure in order to capture any residual autocorrelation is arrived at. Second, the existence of a cointegrating vector, describing the long-run relationship between spot and futures prices, is investigated in this well-specified VECM using the maximum and trace tests proposed by Johansen (1988). Third, once the necessary condition for unbiasedness, that of the existence of a cointegrating relationship, has been identified, the unbiasedness hypothesis is investigated by testing parameter restrictions on the cointegrating vector using the LR statistic of eq. (3).

Regarding the first step, a combination of model selection criteria (SBIC, 1978, and AIC, 1973) and LR tests (see footnote 6) is used to determine the lag length and the

specification of the deterministic components in the VECM. Lag lengths of 2, 1, and 3 are selected for the one-, two-, and three-months data, respectively, and the deterministic components include an intercept in the cointegrating vector in all cases.12 Residual diagnostics on the selected models for first and higher order serial correlation, normality, and ARCH effects show that the models are well specified.

The second step in testing the unbiasedness hypothesis involves determining the existence of a cointegrating relationship between spot and futures prices using the maximum and trace tests. The estimated statistics in Table III, indicate that spot and futures prices are cointegrated. However, Johansen's tests are biased toward finding cointegration too often in finite samples. In particular, Cheung and Lai (1993) find that the finitesample bias of Johansen's tests is a positive function of T/(T - kp), where k is the number of variables in the VECM. Reimers (1992) suggests a small-sample correction of the test statistics for the cointegrating rank. This correction is found to improve the properties of the cointegration tests, particularly in moderately sized samples, and consists of using the factor (T - kp) instead of T in the computation of the [lambda]^sub max^ and [lambda]^sub trace^ tests. Use of the adjusted statistics, denoted as [lambda]^sup *^^sub max^ and [lambda]^sup *^^sub trace^, confirms that spot and futures prices one, two, and three months prior to maturity are cointegrated.

The unbiasedness hypothesis is examined next by testing the restrictions [beta]^sub 1^ = 0 and [beta]^sub 2^ = - 1 in the cointegrating relationship [beta]'X^sub t-1^ = (1 [beta]^sub 1^ [beta]^sub 2^)(S^sub t-1^ 1 F^sub t-1;t-n-1^)'. If these restrictions hold, then the price of a futures contract is an unbiased predictor of the realized spot price. The estimated coefficients of the cointegrating vectors, along with the residual diagnostics for the models, are presented in Tables IV, V and VI for the monthly, two-months and quarterly data respectively. For the one and two-months futures prices, the null hypothesis of unbiasedness cannot be rejected at conventional levels of significance.13 However, for the quarterly futures prices, the restriction is rejected at the 5% level.14 In order to check whether the rejection of the joint hypothesis is driven by the presence of a significant intercept term or from the coefficient of the futures price being significantly different from one we test individually for the null hypotheses b1 4 0 and b2 4 11; our results indicate that b1 and b2 are individually significantly different from 0 and 1, respectively. Hence, we conclude that futures prices one and two months prior to maturity are unbiased expectations forecasts of the realized spot prices; on the other hand, futures prices three months from maturity provide biased forecasts of the realized spot prices.

The issue that arises is to pinpoint what could possibly create biases on futures prices three months from maturity. As already emphasized, the market is characterised by low trading volume and most of the trading concentrates in contracts that are near to maturity. Gilbert (1986) argues that thin trading in a futures market creates worries about execution of trading orders because attempts to trade at quoted prices may change these prices; this may generally result in some forward market bias. Hence it seems that

limited liquidity divorces futures prices from being unbiased forecasts of the realized spot prices three months from maturity. On the other hand, as the contract approaches its maturity day, trading activity in the market increases; this forces futures prices one and two months from maturity to reflect more accurately the expected spot prices on the maturity day of the contract. This finding is in contrast to the findings of Antoniou and Holmes (1996), in the examination of the FTSE-100 contract, who argue that increased trading activity in the market creates price biases in futures prices. It seems that the opposite is true for the freight futures market.

Another possibility is that the bias in futures prices three months prior to maturity reflects imbalances between long and short hedging demand in the market. BIFFEX trades the expected value of a service, which is essentially a nonstorable commodity. Pricing biases seem to be more prevalent for the markets of nonstorable commodities. Empirical studies by Kolb (1992) and Deaves and Krinsky (1995) indicate the existence of significant positive returns in the futures prices of three nonstorable commodities (feeder cattle, live cattle, and live hogs); moreover, these returns increase as the time to maturity of a futures contract increases. Conditional on the assumption of rational expectations, these findings are consistent with the theory of normal backwardation, advanced by Keynes (1930), which hypothesizes that hedging pressures create a differential between futures prices and expected spot prices at contract expiration (that is, a risk premium). The significance of the hedging forces as a factor linking futures and spot prices for nonstorable commodities is also emphasized by Gray and Tomek (1970), who argue that, for these markets, futures prices are linked to expected spot prices through the balance of hedging forces; if these forces represent only one type of futures market position (either net long or short), then a price bias may result.

To investigate the effect that this bias has on the short-run properties of spot and futures prices, we examine the estimated error correction coefficients, [alpha]^sub 1^ and [alpha]^sub 2^, presented in Tables IV, V, and VI. For the monthly and two-months price series, the estimated error correction coefficient on the futures price, [alpha]^sub 2^, is positive and statistically significant, whereas the coefficient on the spot price, [alpha]^sub 1^, is statistically insignificant. The sign of the coefficient on the futures price is in accordance with convergence toward a long-run equilibrium relationship; that is, in response to a positive forecast error at period t - 1 (that is, S^sub t-1^ > F^sub t-1;t-2^), the price of the futures in the next period will increase in value thus restoring the long-run equilibrium. This finding is consistent with the hypothesis that past forecast errors affect the current forecasts of the realized spot prices-that is, the futures prices, but not the spot prices themselves. Therefore, only the futures price responds to the previous period's deviations from the long-run equilibrium relationship and does all the correction to eliminate this disequilibrium.

Turning into the quarterly price series, we can see that both [alpha]^sub 1^ and [alpha]^sub 2^ are positive and significant at the 5% level. This is consistent with our empirical findings regarding the presence of a bias in the quarterly futures prices. If

futures prices provide unbiased forecasts of realized spot prices, they should contain all the information that is relevant in forecasting future spot prices. The existence of a systematic bias on the other hand implies that past forecast errors affect the realized spot prices. The positive signs of both error correction coefficients indicate that a positive forecast error at period t - 1 will force both the futures and the spot prices to increase. Hence any disequilibrium at period t11 is also carried forward to period t, as would be expected by the existence of a bias in futures prices.

The issue that arises is whether the bias on the quarterly futures prices represents a (possibly time-varying) risk premium, conditional on the assumption of rational expectations. A common method of characterizing time-varying risk premia in different futures and forward markets is through the family of autoregressive conditional heteroskedasticity (ARCH) models15; for instance, Domowitz and Hakkio (1985) model the risk premia in the foreign exchange market using the ARCH in mean model, whereas Hall and Taylor (1989) investigate the existence of risk premia in the tin, lead, copper, and zinc contracts at the LME using the GARCH in mean model.

In order to identify the behavior of the bias in the BIFFEX prices three months from maturity, we specify the conditional mean of the threemonth forecast error (that is, S^sub t^-F^sub t;t-3^) using Box-Jenkins (1976) techniques. Our results, presented in Table VII, indicate that the forecast error follows a moving average (MA) process of order 1. However, modeling the conditional variance of the forecast error as a GARCH process does not yield any significant results16; this is consistent, as well, with the residual diagnostics from the estimated model that do not indicate the existence of ARCH effects. Therefore we conclude that the bias in futures prices three months from maturity is a function of the previous period's error term. This can be interpreted either as evidence for the existence of a risk premium that follows an MA(1) process or as the result of market irrationality because the market fails to embody in the current futures prices a systematic time series component of the forecast error (see also in Copeland, 1991). Unfortunately, we cannot distinguish between these two alternatives, because any test of the unbiasedness hypothesis is a joint test that there is no risk premium and that market agents are endowed with rational expectations (Fama, 1991). Therefore failure to accept the unbiasedness hypothesis can be attributed to failure of either of these two reasons. Nevertheless we can still investigate whether the existence of a systematic bias in futures prices (which implies that futures prices do not represent accurate predictors of the realized spot prices) affects the price discovery function of the market.

FORECASTING PERFORMANCE OF FUTURES PRICES

This section investigates the ability of end-of-month futures prices to predict the realized spot prices on the maturity day of the contract. The forecasting performance of futures prices is an issue that has been investigated in the literature for different futures markets. For instance, Ma (1989) and Kumar (1992) compare the forecasting accuracy

of oil futures prices to forecasts generated from time-series and random walk models, whereas Hafer et al. (1992) compare the forecasting performance of Treasury-bill futures prices to that of forward prices and survey data. Broadly speaking, it is found that futures prices provide superior forecasts of the realized spot prices than do forecasts generated from alternative models, although their forecasting performance diminishes as the forecast horizon increases.

The purpose of the forecasting exercise is to investigate whether one can obtain more accurate forecasts of the settlement prices one, two, and three months ahead by employing time-series models rather than using the readily available information provided by futures prices. Futures price forecasts are compared to forecasts generated by bivariate VECM and univariate ARIMA and Holt-Winters (Holt, 1957, and Winters, 1960) exponential smoothing models. The performance of these models in forecasting the BFI has been investigated by other studies in the literature. Cullinane (1992) finds that ARIMA models provide the most accurate forecasts of the BFI for a forecast horizon up to seven days, whereas, for greater lead times, the Holt-Winters model provides superior forecasts. On the other hand, Kavussanos and Nomikos (1998b) find that a VECM of spot and futures prices outperforms these model specifications for forecasts up to 15 days ahead. For comparison purposes, we also consider a benchmark random walk model. This model assumes that BFI prices at time t - n are the most accurate predictors of settlement prices at time t, S^sub t^; therefore it uses information from the historical spot prices to generate forecasts of the future settlement prices and requires no estimation.

In order to compare the forecasting performance of futures prices with that of time-series models, care must be taken to ensure that forecasts from the latter correspond precisely to the forecasts implied by the futures prices. BIFFEX prices converge to the settlement price at the maturity day of the contract and hence, the futures price n months from maturity provides a forecast of the settlement price for this particular day. Because the settlement price of the futures contract is calculated as the average of the BFI over the last five trading days of the contract, timeseries models should be estimated in such a way so that forecasts for these particular days can be obtained. This requirement, in turn, implies that the time-series models should be estimated using daily price data.

Therefore, a VECM is estimated using the most recent 300 daily spot and futures prices of the contract that is closer to maturity.17 Over the period 1988:04 to 1997:03, 107 different VECM models are identi fied and estimated along the lines described in section 3 of the article.18 Similarly, ARIMA models are estimated for each forecast period utilising at any given point the most recent 300 daily BFI observations. The most parsimonious ARIMA model is identified using Box-Jenkins (1976) techniques and the five general model specifications estimated for each forecast period, are presented in Table VIII. For instance, over the period April 1988 to September 1989, the model that provides the best fit for the BFI series is a first difference model with autoregressive

terms at lags 1 and 3. Finally, using the same set of observations as in the ARIMA modelling procedure, the Holt-Winters model is estimated to generate forecasts for the settlement prices 1, 2, and 3 months ahead.19

The forecasting accuracy of each method is assessed using the following criteria. Mean absolute error (MAE), which measures the absolute deviation of the predicted value from the realized value; root mean square error (RMSE), which attaches a higher weight to larger forecast errors; and, finally, Theil's (1966) inequality coefficient, which takes into account the ability of each method to forecast trends and changes (see Pindyck & Rubinfeld, 1991, for the properties of these forecast criteria). The results in TABLE IX indicate that, apart from the one-month horizon where the VECM provides marginally better forecasts than the futures prices, for the remaining horizons futures prices outperform the other models considered. Given the additional time and effort incurred to estimate the VECM, in comparison to the futures forecasts that are readily available in the market, the results for the one-month horizon do not provide reliable evidence to undermine the apparently better forecasting ability of futures prices.20 We can also note that the forecasting performance of futures prices diminishes as the forecast horizon increases; this is consistent with the findings of Ma (1989) and Kumar (1992) and re flects that more information, regarding the future course of spot prices, is available to market participants when the forecasts are made for a shorter horizon. However, although futures prices display this forecasting weakness, they still provide the best forecasts.

Regarding the other forecasting methods, the VECM outperforms the remaining time-series models for all the forecast horizons; ARIMA models outperform the random walk model for the one- and two-months forecasts-although this is reversed for the three-months. On the other hand, the Holt-Winters model has the worst forecasting accuracy over all forecast horizons. This poor performance can be attributed to the stochastic properties of the BFI series. The Holt-Winters model is equivalent to an ARIMA(0,2,2) model specification; if the underlying series can be identified by a different class of ARIMA models, as is the case in our analysis, then forecasts generated by the Holt-Winters model will be far from accurate (see also Harisson, 1967, and Granger & Newbold, 1986). These results indicate that users of the BIFFEX market receive accurate signals from the futures prices regarding the future course of cash prices.

SUMMARY AND CONCLUSIONS

The unbiased expectations hypothesis suggests that the price of a futures contract before maturity should be an unbiased predictor of the spot price on the maturity day of the contract. Several studies have investigated this hypothesis in different futures and forward markets with mixed evidence. Despite this plethora of studies, there has been no evidence for the unbiasedness hypothesis on a futures market that trades the expected value of a service and is characterised by low trading activity-such as the

freight futures market. Parameter restriction tests on the cointegrating relationship between spot and futures prices indicate that futures prices one and two months from maturity provide unbiased forecasts of the realized spot prices. On the other hand, futures prices three months from maturity are biased estimates of the realized spot prices. However, despite the existence of a bias, futures prices provide more accurate forecasts of the realized spot prices than forecasts generated from VECM, random walk, ARIMA, and the Holt-Winters models. This finding emphasizes the significance of the freight futures market as a price discovery center where information about future supply and demand conditions is assimilated and interpreted in an efficient manner.

The implications of these findings are the following. First, market participants receive accurate signals from futures prices and can use the information generated by these prices in order to guide their physical market decisions; therefore charterers or shipowners can use the futures prices as indicators of the future course of BFI prices. Moreover, for shorter maturities (that is, one and two months), the "average" hedger can use the market efficiently without paying any risk premium. For longer maturities, however, the existence of a bias in futures prices increases the cost of hedging, and it seems that hedgers will have to trade off the relatively higher hedging cost against the risk of being unhedged.

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[Footnote]
1For instance, when futures prices are above the expected spot prices, long hedgers buy the futures contracts at a premium over the price they expect to prevail on maturity.

[Footnote]
2See Table I for this, and the "Description of the Contract and Properties of the Data" section of this article for further discussion.
3For instance, on 1 July 1997, the open interest for the January 1998 contract was 122 contracts, and no contract was open for the months beyond.

[Footnote]
4A stochastic process, S^sub t^, is stationary if: (i) It has a constant mean for all values of t, that is, E(S^sub t^) 4 constant, ...t. (ii) It has a constant variance for all values of t, that is, Var(S^sub t^) = constant, ...t. (iii) It has a constant covariance for all values of t, that is, Cov(S^sub t^, S^sub t+n^)=constant, ...t. If a stochastic series must be differenced once in order to become stationary, the series contains 1 unit root and is said to be integrated of order 1, denoted as I(1). If S^sub t^ and F^sub t;t-n^ are I(1) series, any linear combination among these two series will also be I(1). However, there may be a number [beta] such that S^sub t^ - [beta]F^sub t;t-n^ = [epsilon]^sub t^ is stationary. In this special case, Engle and Granger (1987) define the series S^sub t^ and F^sub t;t-n^ as cointegrated of order (1,1) [denoted as CI(1,1)], and the regression S^sub t^ - [beta]F^sub t;t-n^ = [epsilon]^sub t^ is called the cointegrating or equilibrium regression.

[Footnote]

5Because the rank of Pequals the number of its nonzero eigenvalues, the number of the cointegrating vectors can be obtained by estimating the number of these eigenvalues that are significantly different from zero using the following statistics proposed by Johansen (1988); [lambda]^sub trace^(r) = ... and [lambda]max(r,r + 1)= - T ln(1 - ...^sub r+1^) where T is the number of usable observations and ... are the estimated eigenvalues of the [Pi] matrix. [lambda]^sub trace^ tests the null that there are at most r cointegrating vectors against the alternative that the number of cointegrating vectors is greater than r. [lambda]^sub max^ tests the null that the number of cointegrating vectors is r against the alternative of r + 1.

6Johansen (1991) proposes the following statistic to test for the appropriateness of including an intercept term in the cointegrating vector against the alternative that there are linear trends in the level of the series;

...

where ... and ... represent the smallest eigenvalues of the model that includes an intercept term in the cointegrating vector and an intercept term in the short run model respectively. Acceptance of the null hypothesis indicates that the VECM in eq. (2) should be estimated with an intercept term in the cointegrating vector.

[Footnote]

7For the period 1985 to June 1988 only 4 contract months were traded in the freight futures market; January, April, July, and October (the quarterlies). From July 1988 "spot" and "prompt" months were introduced. In other words, since July 1988, at any given month the following contracts were traded; the current month, the following month, and January, April, July, and October up to 18 months ahead. This trading pattern was altered again in October 1991 when trading in a second "prompt" month was introduced. Therefore, from October 91 the following contracts are always traded: the current month, the following two months and January, April, July, and October up to 18 months ahead. For instance, on 30 April 1997 contracts for delivery in the following months were traded: April, May, June, July, October, January 98, April 98, and July 98.

8The use of the settlement price of the futures contract (that is, the average of the BFI over the last five trading days of the contract), rather than the BFI price on the maturity day, was suggested by a referee. The empirical results from using the BFI prices as the realized spot prices are qualitatively the same to the ones reported in this article.

9The first observation coincides with the introduction of the "spot" and "prompt" month contracts. Thus futures forecasts one month ahead are sampled every month.

10In contrast, for the period prior to October 1991, futures forecasts two months ahead are available every three months.

[Footnote]

11These results are available from the authors on request.

12This refers to the lag length of an unrestricted VAR in levels as follows; X^sub t^ = [Sigma]^sup P^^sub i-1^ A^sub 1^X^sub t-1^ + V^sub t^. A VAR with p lags of the dependent variable can be reparameterised in a VECM with p-1 lags of first differences of the dependent variable plus the levels terms.

[Footnote]

13As additional supporting evidence for our results on the two-months prices, inference is also carried out using the Phillips and Hansen (1990) fully modified least squares (FM-OLS) estimator. Use of this method is motivated by the overlapping observations problem, present in the two-months price series (Hansen & Hodrick, 1980). In the presence of serial correlation induced by the overlapping futures forecasts, Moore and Cullen (1995) assert that use of the FM-OLS estimator is more appropriate than Johansen's MLE. In line with Johansen's tests, results from the FM-OLS estimation are supportive of the unbiasedness hypothesis in the two-months futures. Results from these tests are available from the authors on request.

14As suggested by a referee, the rejection of the unbiasedness hypothesis in the quarterly prices may be attributed to the small number of usable observations employed in the analysis (T = 33). As in the case of the trace and max eigenvalues tests, the LR tests for linear parameter restrictions on the cointegrating relationship may reject the null hypothesis too often in small samples. Psaradakis (1994) suggests correcting the LR test of eq. (3) by a factor of (T - m/k)/T where m is the number of estimated parameters in the VECM subject to the reduced rank restriction P 4 ab. For the quarterly prices test, this correction yields a v2(2) statistic of 6.82, with a 95% critical value of 5.99, which still rejects the null hypothesis of unbiasedness.

[Footnote]

15For an overview on the developments in the formulation of ARCH models and a survey on their empirical applications in finance, see Bollerslev et al. (1994).

16These results are available from the authors on request.

[Footnote]
17In line with this article, Ma (1989) and Kumar (1992) also estimate models from daily data using a rolling regressions procedure.
18For instance, the first model is estimated using 300 daily observations of spot and futures prices up to 29 April 1988, which is the last trading day of the July 1988 contract three-months before expiry. The estimation procedure is as follows; Dickey-Fuller (1981) and Phillips-Perron (1988) tests are performed on the spot and futures prices to identify their order of integration and the existence of a stationary relationship between them is investigated using the Johansen (1988) tests. The resulting VECM produces forecasts of the BFI prices for the last five trading days of the July 1988 contract. The arithmetic average of these forecasts yields the forecasted value of the settlement price. This estimation procedure is repeated for the subsequent periods, utilizing at any given point in time the most recent 300 daily spot and futures observations.
19For a review of different classes of exponential smoothing models see Gardner (1985).

[Footnote]
20The forecasting performance of futures prices is also investigated using the BFI prices, rather than the settlement prices, at the expiration day of the contract. Results from these tests also indicate that futures prices provide superior forecasts than the alternative model specifications.

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